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Market versus Administrative Reallocation of Agricultural Land in a Period of Rapid Industrialization

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Property rights in China are moving in two different directions. In some villages, private rights are secure and to some degree marketable; in other villages, individual rights are increasingly restricted and subject to more regulation and reallocation. Administrative reallocation tends to promote more equal access to land, but the price paid for the social insurance of land tenure may be forgone investment.

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Summary findings

Under communal farm production, there was little incentive to work hard: the communal system guaranteed a livelihood, and there were few private gains from additional efforts. The reform that introduced the household responsibility system in China in the early 1980s sharpened individual work incentives by assigning specific plots and the rights to residual income to individual households.

However, the household responsibility system left unresolved questions about the reallocation of land over time — questions that have become increasingly important (for both efficiency and equity) with the rapid growth of the nonfarm economy.

Carter and Yao use household and village data to show that the initially egalitarian distribution of land is becoming more dispersed over time.

In what has become a hybrid property rights system, in some areas local village leaders (the cadre) were empowered to periodically redistribute land between households on the basis of economic and demographic changes among households. In other villages, households were granted much greater immunity against redistribution of any sort.

Similarly, villages differed in the degree to which individual households could trade land among themselves. Some villages did not regulate the practice, and others required village approval or prohibited land rental relationships.

Carter and Yao use simulated maximum likelihood methods to estimate hybrid panel models of the determinants of both market-based and administrative reallocation of land. They also use them to estimate the insecurity-induced investment costs of market-based reallocation of land.

They find that administrative reallocation responds to the increasing inequality but nonmarket reallocations come at a significant cost in forgone investment.

This paper — a product of Rural Development, Development Research Group — is part of a larger effort in the group to study the determinants and impact of property rights systems and land tenure regimes in the process of development. Copies of the paper are available free from the World Bank, 1818 H Street NW, Washington, DC 20433. Please contact Maria Fernandez, room MC3-542, telephone 202-473-3766, fax 202-522-1151, Internet address mfernandez2@worldbank.org. Policy Research Working Papers are also posted on the Web at http://www.worldbank.org/html/dec/Publications/Workpapers/home.html. The authors may be contacted at carter@aae.wisc.edu or yyao@ccer.pku.edu.cn. October 1999. (35 pages)

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MARKET VERSUS ADMINISTRATIVE REALLOCATION OF AGRICULTURAL LAND IN A PERIOD OF RAPID INDUSTRIALIZATION*

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MARKET VERSUS ADMINISTRATIVE REALLOCATION OF AGRICULTURAL LAND IN A PERIOD OF RAPID INDUSTRIALIZATION

Section 1 Introduction

The legalization and subsequent widespread adoption of the "Household Responsibility System" (HRS) in the early 1980s restored the peasant household as the primary production unit in Chinese agriculture. By assigning specific plots and residual income rights to individual households, the HRS sharpened individual incentives for the provision of careful labor and management. Many authors (e.g., Lin, 1992 and MacMillan, Whalley, and Zhu, 1989) have attributed the subsequent rapid growth of agricultural output to the fact that the HRS broke the syndrome of the iron rice bowl in which little labor effort was provided on the grounds that the prior communal production system guaranteed a livelihood and private gains from additional efforts were sparse.[1]

While decisively individualizing residual income rights, the HRS more generally resulted in a hybrid property rights system in which use, transfer and reallocation rights were variously distributed between the individual land users and their villages that remained the formal custodians of land that legally remained the property of the state. As Liu *et al.* (1998) show, in some areas, local village leaders (the cadre) were empowered to periodically redistribute land between households based on economic and demographic changes among households, while in other villages, households were granted much greater immunity against redistribution of any sort. Similarly, villages differed in the degree to which individual households could trade land among themselves, with some villages not regulating the practice at all, whereas others required village approval or

[1] There has, however, been some dispute about the severity of the labor incentive problems under the communal system and the extent the rapid agricultural of the 1980's was due to the HRS or the price reforms that shortly followed (e.g., see Putterman, 1991).

simply prohibited land rental relationships. Dong (1994) has argued that the redistributive power of the cadre permits land to serve an important social safety net function, an argument ratified by Burgess (1997) who econometrically shows that the (presumably administratively maintained) egalitarian land distribution in China contributes to low levels of undernutrition. Kung (1997) presents evidence that peasants prefer this mixed to a fully privatized, redistribution-proof system.

At the time of the HRS reform, most villages allocated land to households based on the workforce size, with adjustment sometimes made for the number of its dependents (Liu *et al.*, 1998). Given this distributional mechanism, the initial post-HRS distribution of factor endowments—defined as the ratio of contract land to family labor—across households must have been fairly compressed. Moreover, with limited opportunities for off-farm jobs or for urban migration, there must have been relatively little scope for land and labor exchange between households. Even with limited exchange of these factors, and irrespective of the definition of tenure security and transfer rights, the intensity with which land was cultivated must have also been quite similar across households.[2] While tenure insecurity may have dampened incentives for long term, fixed investment under the mixed HRS property rights regime (an issue to be investigated below), the productivity impacts of the secondary features of the HRS (*e.g.*, limited transfer rights) must have initially been quite modest.

However, the rapid growth and industrialization of the Chinese economy since 1980, with its concomitant withdrawal of labor from agriculture, would be expected to put new pressure on the fluidity and efficacy of the institutions that allocate rural land

[2] Graynor and Putterman (199x) note that Chayanovian forces will tend to equalize factor proportions between households even when land is distributed on a need as opposed to a work capacity basis.

and labor. In this new environment, the absence of further administrative reallocation of endowments with no increase in private land or labor transactions, would imply an increasing dispersion in both the factor endowment and the factor intensity distributions. An increased dispersion in the latter distribution would signal a potentially costly (static) allocative inefficiency problem of under- (and over-) utilized land. Moreover, as Carter and Yao (1998) argue, in an environment of rapidly growing off-farm employment, imperfectly transferable rights may discourage investment by households that expect to exit or reduce their participation in agriculture in the near future, creating a sort of dynamic inefficiency problem as well.

Punctuated by a slowdown in agricultural growth, these new challenges re-ignited debate over rural property rights and the efficacy of the HRS and its hybrid property rights regime. In theory, an allocatively more efficient economy could be achieved either by mechanisms that facilitate rental transactions between households in the face of an increasingly disperse endowment distribution; or, by mechanisms that reallocate distribution of land endowments such that no transactions are not required to equalize factor intensity production across production farms. Matching these theoretical possibilities, proposals for further reform in China range from those that look toward a more administrative mechanisms to adjust the endowment distribution, requiring a recentralization of property rights, to those that propose to rely on decentralized, market-mediated decisions and ask for the deepening of individual tenure security and transfer rights needed to make this mechanism work.

While either market or administrative mechanisms could in theory resolve the new challenges posed by rapid industrialization, this paper turns to empirical data to gauge how these mechanisms actually function in practice. Using panel data on a sample of 80 villages drawn from across rural China in 1988 and 1993, this paper tries to first clarify the extent of these new land tenure problems. Section 1 below first looks at the evolving distributions of endowments and factor proportions in production over the 1988 to 1993 time period. We find using a non-parametric Kolmogorov-Smirnov test that we cannot reject the hypothesis that these distributions have changed over time, and parametric maximum likelihood estimates show that the land-labor endowment distribution has indeed become more disperse over time, revealing the uneven impact of industrial employment growth and other demographic shifts. At the same time, we find that the that factor intensity distribution (in rice production) has, if anything, become slightly less disperse over this time period, indicating that either centralized or decentralized mechanisms are at least keeping pace with the challenge of demographic and industrial change.

How then do these two competing mechanisms work and at what cost? Addressing this first issue, Section 3 explores how and how effectively administrative and market-mediated land reallocations respond to evolving opportunities to improve the efficiency or the equity of the rural economy. Simulated maximum likelihood methods are used to estimate a hybrid panel data tobit model of land reallocations and land rentals. Our strongest finding is that administrative reallocations seems to respond primarily to increasing dispersion in endowment ratios (*i.e.*, to equity concerns) and hence can be seen to play a social insurance function.

Section 4 then turns to see if this social insurance is indeed costly, using data on rice-producing households to explore whether or not the tenure insecurity implied by centralized powers of administrative reallocation significantly dampen investment incentives. Using panel data methods to control for the likely correlation between tenure regime and unobserved factors that influence returns to investment, the analysis in this section shows that tenure insecurity significantly reduces investment. Whether or not this price is one worth paying for this insurance is an issue that will require additional research, as discussed in the paper's final and concluding section.

Section 2 Changes in the Factor Endowment and Factor Intensity Distributions, 1988 to 1993

A well-functioning economy that obtains maximal output from the resources that it utilizes will allocate inputs such that their marginal productivities are equal across different uses and production units. Were this not the case, total output could be increased by shifting inputs from lower to higher productivity locations. Under fairly general assumptions about the nature of technology, dispersion in the ratio of labor allocated to production per-unit area cultivated is a good indicator of the degree to which an agricultural economy has succeeded in equating factor productivities across farm units. While it is unlikely that any real world economy will perfectly equalize factor intensity across households,[3] panel data from a sample of households that were surveyed in 1989 and again in 1994, permits us to explore whether the dispersion in factor proportions has increased over time. These same data also provide a window through

[3] In measurement sense have trouble with unmeasured factor quality differentials that may make it efficient to have different factor proportions. Also expect some time of adjustment such that at any moment of time be some variation.

which to view how the endowment distribution (defined as the ratio of household contract land per-family member) has evolved under hybrid property rights and the pressure of rapid industrialization.

The overall survey covered 800 households spread across 80 villages drawn from 4 provinces (Henan, Jiangxi, Jilin, and Zhejiang).[4] Table 1 gives some descriptive indicators drawn from the approximately 400 surveyed households that primarily cultivate rice. As can be seen, as measured by the coefficient of variation, both the endowment and factor proportions distribution became more disperse for the sample as a whole, with the coefficients of variations for these two distributions rising from 74% to 100%, and 46% to 50%, respectively. Some of this increased variation or dispersion in these distributions may well result from differences between villages. For example, we might expect that more rapidly growing regions will experience an increase in wage rates and a matching fall in the labor intensity of rice production. In order to distinguish this source of increased variation from that occurs because of, say, faulty allocative institutions within the local village economy, Table 1 also presents an intra-village coefficients of variation for the endowment distribution defined as:

$$\sigma_{it}^{\tau} = \left(\sum_{h=1}^{n_{it}} (\tau_{iht} - \bar{\tau}_{it})^2 / n_i\right) / \bar{\tau}_{it} \tag{2-1}$$

[4] The data used in this study come from two comprehensive surveys administered in eight provinces of China in 1988 and 1993 for studies of land tenure evolution and its implications to agricultural productivity. Both surveys have a household and a village questionnaire. Villages and households were chosen from the Rural Survey Base maintained by the Rural Survey Team of the State Statistical Bureau of China. A household questionnaire was administered to ten randomly selected households in each village and asked questions ranging from land rights, land transactions and annual agricultural production to off-farm employment. The village questionnaire asked questions about land tenure arrangements at the village level both in the surveyed years and in the history. In order to eliminate extraneous noise, the analysis of investment and changing dispersion in factor proportions in production relies only on the subset of villages where rice is the predominant crop.

where τ_{hit} is the contract land-family labor ratio for household h in village i in year t; n_{it} is the number of sampled households in village i in year t; and, the mean endowment ratio for village i in year t is $\bar{\tau}_{it} = \sum_{h=1}^{n_i} \tau_{iht} / n_{it}$. Letting ℓ denote the ratio of land to labor in rice production, a similar expression defines the intra-village coefficient of variation in factor proportions, σ_{it}^{ℓ}. As can be seen, the intra-village coefficient of variation of factor endowments rose from 43% to 54%, whereas the intra-village the same measure for the distribution of factor proportions actually declined slightly (34% from to 32%).

The Kolmogorov-Smirnov statistics reported in Table 1 test the hypothesis that the 1988 and 1993 distributions are statistically different from another—*i.e.*, they test whether or not the difference in the distributions signaled by the coefficients of variation are large enough that we can reject the hypothesis that the 1988 and 1993 samples were drawn from the same population distributions.[5] Table 1 reports the value of the Kolmogorov-Smirnov test statistic and the figure in parenthesis gives the probability of randomly generating a test statistic of that size under the maintained hypotheis that the two distributions are identical. At a 5% significance level the data thus reject the hypotheses that the 1988 distributions are identical to the 1993 distributions for both endowments and factor proportions. However, we cannot reject the hypothesis that the intra-village factor proportions distribution was the same in 1988 as it was in 1993. There is, however, stronger evidence that intra-village distribution of factor endowments

[5] The distribution free Kolmogorov-Smirnov test compares the empirical cumulative density functions for two distributions, asking if the largest difference that occurs between the two functions is so large that it is unlikely to have happened based on random draws from identical distribution.

has changed as we can reject the hypothesis of identical distributions at the 11% significance level.

While these tests confirm which distributions have changed in a significant way, they provide no information on the nature of the change nor its efficiency implications. In order to gain a more precise (and visual) idea about the nature of these evolving distributions, we assume that the factor endowments and factor proportions are distributed according to a gamma distribution:

$$f(\tau_{ht}) \sim \Gamma(\alpha_t^{\tau}, \beta_t^{\tau})$$

(2-2)

$$f(\ell_{ht}) \sim \Gamma(\alpha_t^{\ell}, \beta_t^{\ell})$$

where $f(\bullet)$ is the probability distribution function for the distribution of factor endowments, τ_{it}, and factor intensity, ℓ_{it} for household h in period t. The gamma distribution has been chosen for this analysis because it is a flexible distribution form that can take on a wide variety of shapes, from normal to skewed to exponential, depending on the value of the parameters of the distribution. In addition, in order to explore differences between the 1988 and 1993 distributions, we specify the two parameters of the distribution to be linear functions of time:

$$\begin{aligned} \alpha_t^j &= \alpha_o^j + \alpha^j D_t \\ \beta_t^j &= \beta_o^j + \beta^j D_t \end{aligned} \tag{2-3}$$

where $j = \tau, \ell$ and the binary variable D_t takes on the value of one for 1993. Table 1 presents maximum likelihood estimates of the parameters defined by (2-2) and (2-3), while Figure 1 displays the fitted or estimated probability functions.

As can be seen in Table 1, the ML estimates indicate that parameters of both distributions have changed over time in a statistically significant fashion. Figure 1a, which graphs the estimated endowment distributions for 1988 and 1993, shows that over time the endowment distribution has indeed become more disperse, with greater density in both the upper and lower tails of the estimated distribution.

Figure 1b displays the estimated distribution of the labor to land (or factor proportions) ratio for the sample of rice producers. Already by 1988, the distribution of factor intensity appears relatively disperse. The mode for the estimated 1988 factor proportions distribution is about 200 hours of labor per-*mu*. However, as can also be seen, there is a not inconsequential number of farms that apply as few as 100 hours of labor per-*mu*, acting as if labor is very scarce or dear. Similarly, there are a number of units that behave in the opposite way, allocating as much as 400 hours of labor per-*mu* of cultivated rice. Either reallocation of land from the land abundant to the land scarce farms, or reallocation of labor from labor abundant to labor scarce farms would be expected to increase total agricultural output from the same resource base. Assuming constant returns to scale, an output labor elasticity of one third and equal use of non-labor inputs per-*mu* on all farms, this estimated dispersion in labor intensity implies a 2.2% yield loss (that is, yields would average 2.2% higher if all farm units produced at the average labor intensity of 240 hours per-*mu*).

Compared to the 1988 distribution, the 1993 distribution has shifted right. While the upper tail of the distribution has increased (indicating that there is a greater proportion of farms are cultivating their land more intensively in 1993), the lower tail has actually shrunk over the 1988 to 1993 period. Calculated under the same assumptions described in the prior paragraph, the yield loss implied by this 1993 distribution of labor intensity is 2.4%, slightly higher than the 1988 value. Note that because average labor intensity was higher in 1993 than in 1988 (309 versus 240 hours per-*mu*), the physical losses implied by this labor intensity dispersion are greater than the 0.2% yield loss difference would imply.

To summarize the findings of this section, we see that as expected under the pressure of an industrializing economy and ongoing demographic change, the overall endowment distribution has spread out over time, both across and within villages. However, the impacts on distribution of labor intensity in production have been much more muted. While the overall level of allocative inefficiency in the rural, rice-producing economy seems to have increased slightly (to about a 2.4% yield loss), most of that change seems to have occurred between rather than within villages. There is thus some evidence that local institutions and mechanisms are coping relatively well with the challenges of continuing industrialization and labor outmigration.

Section 3 The Responsiveness of Market and Administrative Reallocations to Social and Economic Forces

As described in the introduction above, the endowment and factor proportions distributions in rural China have been evolving under a variety of locally determined institutional mechanisms. While considerable controversy surrounds the relative merits of decentralized, market-based mechanisms versus centralized administrative mechanisms, empirical analysis of how either functions is scarce. An important exception is the study of Turner, Brandt and Rozelle (1998) who estimate the determinants of administrative land reallocation in China. Their theoretical ruminations suggest that cadre (and the village political process) may reallocate land for a variety of reasons, ranging from equity goals to concerns to maximize the productivity of village resources (which they term the "rational village" hypothesis). In the analysis here, we will follow their theoretical lead and test for the significance of both equity and efficiency factors in motivating administrative reallocations.

While finding evidence in favor of what they term the "rational village" hypothesis, Turner *et al.* (1998) also note that market and administrative mechanisms are alternative means for achieving the same goal. In a simple OLS regression, they find that a one unit increase in the amount of land reallocated through the market results in almost unit reduction in the amount of land administratively reallocated. However, they note that this result is not to be trusted given that administrative and market-based reallocations are simultaneously determined, and that any effort to explore their determinants must properly account for this simultaneously. Unfortunately, within the confines of their available cross-sectional data, they confront an intractable identification problem that prevents them from identifying the determinants of both market and administrative reallocations.

3.1 A Switching Regressions Model of Administrative and Market-Mediated Land Reallocation

The panel data available here suggests a resolution to this identification problem and opens the way to an exploration of the responsiveness of market and administrative reallocation to socio-economic forces. Following the theoretical suggestion of Turner *et* *al.* (1998), we begin by noting that the percent of land that the village cadre would desire to administratively reallocated in village *i* PRIOR to the commencement of agricultural production in year *t* (denoted A^{*}_{it-}) may respond either to inequitable intra-village dispersion in the endowment distribution, or to inefficient intra-village dispersion in the factor proportions distribution that were observed to have occurred in the prior production period, *t-1*. Dispersion in the former distribution would signal the potential for equity gains from reallocation, while dispersion in the latter would indicate that reallocation would create efficiency gains. The intra-village coefficients of variations defined in Section 2 above suggests themselves as natural measures of this dispersion. In addition, the percent of village grain output that must be delivered to the state quota system is likely to signal the severity of the pressure upon village cadres to attend to productivity-dampening inefficiency.

The percent of village land that villagers would desire to transact through rental contracts in period *t* (R^{*}_{it}) take place AFTER any administrative reallocations and thus responds to residual opportunities for mutually beneficial exchange that are created by intra-village dispersion in the contemporaneous distributions of endowments and off-farm, wage labor opportunity, where the latter is denoted σ^{w}_{it}.[6]

Table 3 presents descriptive statistical indicators of these dispersion measures, as well as measures of the amounts of land reallocated by both administrative and market-

[6] In one of the few studies that focuses on private rental transactions in China, Carter and Yao (1998) use standard household models to show that households become more likely to undertake land rental transactions as their land:labor factor endowment ratio becomes more extreme and as the household enjoys specific factors and skills that encourage it to specialize in either agricultural or non-agricultural activities. The intra-village dispersion measures hypothesized here to influence the activeness of the village land market capture in a simple way the forces identified by Carter and Yao's model.

based mechanisms. To calculate both these latter two measures, we first estimated the total land stock in a village using the information on household land allocation together with reports of land set aside by the village for future redistributions. The reported rental variables were calculated by summing up the absolute value of all rental transactions (both land rented in and land rented out) and dividing by the estimated village land stock. The administrative reallocation variable is the sum of all land reallocations (positive or negative) reported by surveyed households over the 1988 through 1992 period.[7] As can be seen, on average just under 10% of the village land base was affected by rental transactions in both 1988 and 1993.[8] The total amount of land administratively reallocated over the 1988 to 1992 period was about 11% of the village land base. Unfortunately, we do not have data on the administrative allocations that took place in the wake of the 1993 cropping season.

Consistent with the data from the rice-producing villages analyzed above, the intra-village coefficients of variation measures show on average an increase in the variation in the endowment distribution, and a decrease in the dispersion of the factor proportions distribution.[9] Intra-village variation in off-farm income (as measured by the intra-village coefficient of variation in household non-farm earnings) increased substantially over the 1988 to 1993 period. In 1988, intra-village variation was 93% of mean household off-farm earnings. By 1993, this figure had risen slightly to 98%,

[7] A multi-year measure was used on the grounds that administrative reallocations are costly and therefore take place only periodically. A single year measure would thus be likely to understate the frequency and amount of administrative reallocations.
[8] As measured in the survey, land rentals include both transactions between individuals and temporary lease transactions between villages and households in those villages that did not initially distribute all land to households.
[9] Factor proportions was here measured as the total household input of labor into agriculture divided by the household's cultivated land area.

although the intra-village standard deviation of off-farm earnings more than tripled over this time period (reflecting an increase in average household off-farm earnings from 1350 to 4000 yuan). The increase in both figures reflects the rapid growth of the non-farm economy and the modest increase in household specialization between agricultural and non-agricultural sectors.

While the dispersion measures reflect opportunities for gains from administrative reallocation as well as gains from rental trades, the desired levels of these transactions would be expected to depend on the permissiveness of the village-level property rights regime. As Liu *et al.* (1998) analyze in detail, in some villages the power of the cadre to administratively reallocate land is highly encumbered by regulations that grant individuals considerable tenure security and immunity against reallocation. There is also significant variation across villages in the degree to which individuals are permitted to engage in rental transactions. For the analysis here, we have defined two binary property rights variables, D^A_{it}, and D^R_{it}. The first takes on the value of one when individuals are immune to redistribution, while the second takes the value of one when individuals can rent land out to others with no or minimal restrictions.

In conformity with the preceding observations on the forces that potentially shape market-mediated and administrative reallocations of land, consider the following switching regressions specifications:

$$\begin{aligned} A^*_{it^-} &= z^A_{it-1}\beta^A + D^A_{it}(z^A_{it-1}\delta^A) + \tilde{\varepsilon}^A_{it} \\ R^*_{it} &= z^R_{it}\beta^R + D^R_{it}(z^R_{it}\delta^R) + \tilde{\varepsilon}^R_{it} \end{aligned} \quad , \tag{3-1}$$

where the z are the vectors of explanatory variables described above. Because the property rights regimes are themselves endogenous (*e.g.*, see the discussion in Liu *et al.*, 1998), it is unlikely that either regime switching or the dispersion vectors are orthogonal to the error terms in (3-1).[10] For example, villages close to the "contagion" of rapidly growing privately owned or joint venture factories may tend to have both more individualized property rights and greater intra-village dispersions in endowments, factor use and opportunity. And yet this same contagion may directly induce more risk taking, greater specialization and a more active land rental market.

Over the relatively brief period of our survey, we assume that the troublesome portion of the error terms in the rental equation is time invariant and following the lead of the panel data literature (e.g., Mundlak, 1978) we will assume that we can decompose it as:

(3-2) $$\tilde{\varepsilon}_{it}^{R} = \alpha_{i}^{R} + \varepsilon_{it}^{R}$$

where ε_{it}^{R} is purely white noise, uncorrelated with the D_{it} and z_{it}, and that the troublesome correlation between the error term is given as:

(3-3) $$\alpha_i^R = \bar{z}_i^R \gamma + \omega_i^R$$

where $\bar{z}_i^R = \sum_{t=1}^{2} z_{it}^R / 2$, and the γ are nuisance parameters.[11] In the estimation that follows we will impose the following normality assumptions, $\varepsilon_{it}^R \sim N(0,\sigma_R^2)$ and $\omega_i^R \sim N(0,\sigma_\omega^2)$.

While we can similarly decompose the error for the administrative land equation as:

(3-4) $$\tilde{\varepsilon}_{it}^{A} = \alpha_i^A + \varepsilon_{it}^A,$$

[10] Note that because we use *intra*-village dispersion measures, we already control for many differences between villages that might lead to inter-village dispersion.

[11] Note that the $\underline{\bar{x}}_i^A$ are being used as instruments for the latent component α_i^R.

we would not be able to identify all the nuisance parameters in the analogue to equation (3-3) because we have only a single cross section of observations on administrative reallocations. Instead we suppose that latent, time invariant component in the reallocation equation, α_i^A is a scalar multiple of that in the rental equation:

$$\alpha_{it}^A = \Delta(\alpha_{it}^R) = \Delta(\bar{x}_i^R \gamma + \omega_i^R), \tag{3-5}$$

where Δ is the scaling parameter. Note that under (3-5) we are instrumenting for the time invariant term in the reallocation equation using the village averages for the time invariant term in the market rental equation. We will assume that the white noise error component in (3-4) is normally distributed as $\varepsilon_{it}^A \sim N(0,\sigma_A^2)$. We turn now to develop a method to estimate the system defined by (3-1)-(3-5).

3.2 Simulated Maximum Likelihood Estimation

Both the dependent variables in (3-1) are censored. Administrative reallocations are costly to organize, and presumably do not take place unless there is a significant backlog of equity- and, or efficiency-promoting reallocations. Matching this expectation, in the available data, only 53% villages engaged in any reallocation over the 1988 to 1993 period.

While land rental transactions are by definition more decentralized and presumably less costly to individually initiate, it is also likely that an active land market does not really come into existence until there is some critical minimum number of potential transactions to be made. This observation implies that rental transactions are likely to be censored as well, and indeed the data show that no rental transactions are recorded in 81% of the villages in 1988 and 63% in 1993.

Reflecting this censorship, we can now defined the observable, censored dependent amount of administrative reallocation, A_{it}, as:

$$A_{it^-} = \{ \begin{matrix} A^*_{it^-}, \text{if } A^*_{it^-} > 0 \\ 0, \text{otherwise} \end{matrix} .$$

A similar expression defines observable rental transactions, R_{it}. Unfortunately, in the presence of this censorship, consistent estimation of panel data model defined above cannot be attained using fixed-effect, analysis of covariance estimators.[12]

To get around this problem, we follow Gourieroux and Monfort (1993) and adapt simulated maximum likelihood methods to our problems. Note that without loss of generality, we can rewrite ω_i^R in (3-3) as:

$$\omega_i^R = \theta \upsilon_i^R ,$$

where we assume that $\upsilon_i^R \sim N(0,1)$, and θ is a parameter to be estimated. While we do not actually observe υ_i^R, we can simulate H replications of it using Monte Carlo methods. Conditional on a given simulated value, υ_{ih}^R, the full model to be estimated can be written as:

$$(3\text{-}6) \quad A_{it^-} = \{ \begin{matrix} A^*_{it^-} = z^A_{it-1}\beta^A + D^A_{it}(z^A_{it-1}\delta^A) + \Delta[\bar{z}_i^R\gamma + \theta\upsilon^R_{ih}] + \varepsilon^A_{it}, \ \text{if } A^*_{it^-} > 0 \\ 0, \text{otherwise} \end{matrix}$$

$$(3\text{-}7) \quad R_{it} = \{ \begin{matrix} R^*_{it} = z^R_{it}\beta^R + D^R_{it}(z^R_{it}\delta^R) + [\bar{z}_i^R\gamma + \theta\upsilon^R_{ih}] + \varepsilon^R_{it}, \ \text{if } R^*_{it} > 0 \\ 0, \text{otherwise} \end{matrix}$$

for which the likelihoods conditional on that simulated value are given by:

[12] See for example, the discussion in Hsaio (1986). Intuitively the problem is that differencing time periods no longer necessarily sweeps away the time invariant components as R^*_{it} and A^*_{it} are no longer necessarily observed in both periods.

$$\text{(3-8)}\quad \begin{aligned} L^{A}_{ith} = & \{\prod_{it \ni A_0} \Phi[(-z_{it}^{A}\beta^{A} - D_{it}^{A}(z_{it}^{A}\delta^{A}) - \Delta(\bar{z}_{i}^{R}\gamma + \theta\upsilon_{ih}^{R}))/\sigma_{A}]\} \\ & \{\prod_{it \ni A_1} \phi[(A_{it^-} - z_{it}^{A}\beta^{A} - D_{it}^{R}(z\delta_{it}^{A}) - \Delta(\bar{z}_{i}^{R}\gamma + \theta\upsilon_{ih}^{R}))/\sigma_{A}]\} \end{aligned}$$

and,

$$\text{(3-9)}\quad \begin{aligned} L^{R}_{ith} = & \{\prod_{it \ni R_0} \Phi[(-z_{it}^{R}\underline{\beta}^{R} - D_{it}^{R}(z_{it}^{R}\delta^{R}) - \bar{z}_{i}^{R}\gamma - \theta\upsilon_{ih}^{R})/\sigma_{R}]\} \\ & \{\prod_{it \ni R_1} \phi[(R_{it} - z_{it}^{R}\underline{\beta}^{R} - D_{it}^{R}(z_{it}^{R}\delta^{R}) - \bar{z}_{i}^{R}\gamma - \theta\upsilon_{ih}^{R})/\sigma_{R}]\} \end{aligned}$$

where A_0 is the set of observations with non-positive administrative transactions, A_1 the set with positive administrative transactions, R_0 the set with no market rentals and R_1 the set with positive rentals. The functions $\phi(\bullet)$ and $\Phi(\bullet)$ respectively denote the probability and cumulative density functions for the standard normal distribution.

For H replications of υ_{i}^{R}, we can write the mean simulated log likelihood function for the entire model as:

$$\text{(3-10)}\qquad \overline{L}_{it} \equiv \ln[\frac{1}{H}\sum_{h=1}^{H} L_{ith}^{A}(\varepsilon_{it}^{A} \mid \upsilon_{ih}^{R})] + \ln[\frac{1}{H}\sum_{h=1}^{H} L_{ith}^{R}(\varepsilon_{it}^{R} \mid \upsilon_{ih}^{R})].$$

Maximization of the mean simulated log likelihood will yield consistent estimators of the parameters in the model above, and will yield a good approximation of the true likelihood even with a moderate H, as long as different values of υ_{i}^{R} are drawn for each observation (Gourieroux and Monfort, 1993). In the analysis that follows, H is set equal to 50.

3.3 Results

Table 3 presents the parameter estimates that result from maximizing (3-10) for two slightly different model specifications. In addition to these SML estimates,Table 3 also present estimates derived by simpler tobit methods that ignore the potentially problematic relationship between the latent village effect (α_i) and the included variables. For the administrative reallocation equation, this simpler tobit model reduces to a single cross-sectional estimate, while for the rental equation it results in a simple pooling of the two cross sections.

Like the SML-1 estimates, the simple, pooled data tobit estimates using the intra-village coefficient of variation to measure the dispersion in off-farm income earning opportunities. As the table shows, these simple pooled data tobit estimates more or less correspond to our prior expectations about the forces that motivate land rental transactions. When the rental rights variable $D^R = 1$ (indicating that individuals are free to rent their land out), increasing dispersion of the endowment distribution is estimated to increase the percentage of village land that is reallocated through the rental market. The variables are scaled such that an increase in the coefficient of variation for the endowment ratio by 10 percentage points would increase the percentage of land rented by a hefty 5 percentage points (recall that we have measured land rented as the sum of land rented in and land rented out). Increased dispersion in off-farm earnings has a not quite significant but positive effect on rental transactions, according to these estimates, although there is no significant effect of freer rental rights on this effect.

While these pooled results are sensible, as argued above they are likely biased as they do not control for unobserved factors that may be simultaneously increasing the demand for rental transactions and the demand for less restricted property rights. To the extent that these factors are invariant over the period of our study, then the panel methods proposed above will permit us to identify the actual structural or *ceteris paribus* effect of changes in trade opportunities and property rights on land rental activities. The coefficients on the terms used to instrument for the latent village effect (the $\bar{z}_i^R$) are nearly all significant, indicating that time invariant characteristics explain a large portion of the overall variation in land rental activity. As the results under the column labeled SML-1 show, the panel estimates for the effect of off-farm earning dispersion are similar (positive, but only significant at the 20% level). While the coefficient on the interaction between endowment dispersion and freer property rights is similar to the simpled pooled estimates, the coefficient that gives the impact of endowment dispersion on market rentals irrespective of the property rights regime has become large and significantly negative.

Turning now to the estimates of the percentage of village land reallocated by administrative processes, we find that the simple tobit, cross-sectional estimates do not match at all with prior expectations. Recall that the property rights variable D^A takes the value of "1" when individuals are immune to periodic redistribution, so that we might expect terms involving D^A to carry a negative sign For the most part, the coefficients in this regression are insignificant, although the significant coefficients related to the endowment dispersion terms carry precisely the opposites signs to what would be expected. Recalling that we are using structural and other conditions in 1988 to explain reallocations that took place between 1988 and 1993, it may be that these results reflect administrative reallocations that took place only after a post-1988 change in the property rights regime. It may also reflect the sorts of biases described above.

The SML estimates are designed to alleviate this latter problem. As described earlier, these are not conventional panel estimates inasmuch as the predicted village effect variable from the rental market equation is used to control for the latent effect in the reallocation equation. The parameter Δ that scales the rental village effect is negative, as expected (*i.e.*, latent factors that increase land rentals decrease land that is administratively reallocated). As can be seen in Table 3, the SML-1 estimates of endowments, quotas and property rights now carry the expected sign, but they are statistically insignificant.

The estimates shown as the SML Panel-2 estimates in Table 3 result from a slight specification change. As noted above, the intra-village coefficient of variation in off-farm earnings changed only slight between 1988 and 1993 even though the magnitude of off-farm earnings more than tripled over this period. Arguably, the intra-village standard deviation in off-farm earnings is the better measure of the potential gains from exchange in the land rental market. As can be seen in Table 3, using this latter measure has surprisingly little effect on the market rental equation. However, it has a major effect on the administrative reallocation estimates (note that this change in variable effects the administrative reallocation equation only through its impact on the estimate of the latent village effect). As can be seen, the variables related to both quotas and endowment dispersion become significant (statistically and economically) under this second specification. While these results should be accepted with caution, they imply that village cadre respond to both equity concerns in reallocating land (in the absence of property rights forbidding reallocation, a 10 percentage point increase in the intra-village coefficient of variation for the endowment distribution brings an estimated reallocation of 2.4% of village land), and that they reallocate more frequently when quota delivery pressure is higher. Individual immunity against such redistributions (D^A=1) seems to block further reallocations for equity purposes, though it has no effect on quota-driven reallocations. While both of these results are consistent with a rational village model of reallocation, the coefficient on dispersion in the land-labor factor intensity variable has becomes significant and counterintuitively negative in the SML-2 regression.

To summarize, the statistically conservative panel data approach used in this section has provided uncertain inference at best concerning factors that shape market and administrative reallocation of land. The data in general offer little insight on the factors that increase land rentals.[13] A clearer picture emerges about administrative reallocation, however, with both equity and production-minded (if not precisely allocative efficiency) concerns driving these sorts of reallocations.

Section 4 Costs of Administrative Reallocation: Tenure Insecurity and Investment Disincentives

While the analysis in the prior section found evidence that administrative land reallocations respond to emerging inequality in endowments, a potential cost of this sort of social insurance is that by attenuating the security with which individuals hold their property rights, it may reduce incentives for investments that are attached to land. The fact that insurance (or any other good) is costly does not by itself mean that it is bad public policy to provide it. The more relevant question concerns how costly that insurance is relative to its benefits and to other ways of providing it. As a first step toward answering this question, this section turns to the household data from the rice producing villages in an attempt to determine whether or not the prospect of

[13] These disappointing results suggest that household level analysis is necessary to gain insight on the forces driving rental transactions, such as that in Carter and Yao (1998) discussed in note 7 above.

administrative reallocations dampens fixed investment.

The kinds of fixed investments observed in the data include wells, irrigation and drainage ditches, ponds, application of organic fertilizers whose fertility effects persist for several years, land leveling, soil improvement, and others. The major cost associated with all these investments is the labor hours devoted to producing them. For each of the rice-producing household in the sample, we calculated the total hours devoted to such activities in both 1988 and 1993. Because 19% of the households in the sample did not dedicate any labor to these investment activities, we adopt a tobit approach:

$$(4\text{-}1)\quad k_{jt} = \begin{cases} k_{jt}^{*} = z_{jt}^{k}\beta^{k} + \tilde{\varepsilon}_{jt}^{k}, \text{ if } k_{jt}^{*} > 0 \\ 0, \text{ otherwise} \end{cases}$$

where k_{jt}^{*} is the desired (uncensored) level of fixed investment by household *j* in year *t*. The factors likely to influence desired investment that we include in the vector z^k of explanatory variables include a measure of land quality (% of land that is flat and can be irrigated) a measure of farm fragmentation (the number of parcels held by the household), measures of household endowments (age and education of household head, and household contract land endowment per family member) and measures of the opportunity cost of labor (non-agricultural employment opportunities). Regional and time dummies are also included in z^k to capture the effect of otherwise unmeasured factors. Finally the property rights variable defined earlier, D_{jt}^{A}, that signals individual immunity against administrative land reallocation is included as a measure of tenure security. Table 2 above presents mean values for these explanatory different variables.

A conventional tobit estimator of (4-1) derived by simply pooling the data together form the two survey years is given the first column of Table 5. Aside from the year and county dummy variables, none of the structural factors significantly explain investment, with the exception of the land fragmentation variable. According to the latter estimate, as a farm shifted from only 1 parcels to the sample average of about 10, investmetn would decrease by about s fragmentation (measured as the number of parcels) increases, investment would decrease by about 1.5 hours per *mu* (or 25% of the sample average investment level). Surprisingly, the estimated effect of tenure security is both small and statistically insignificant.

The consistency of these tobit estimates can, however, be contested on the grounds that the degree of tenure security is itself endogenous, and perhaps is itself caused by the same factors that make investment more profitable. In more straightforward econometric terms, tenure security and perhaps other variables in (4-1) may not be orthogonal to latent investment-promoting factors that are subsumed in the error term. Assuming that the latent, investment-promoting household and village effects are invariant over the five year time period covered by the panel data, we can again employ the panel data methods introduced in the prior section and decompose the error term in (4-1) as

$$(4\text{-}2) \qquad \tilde{\varepsilon}_{jt}^{k} = \alpha_j^k + \varepsilon_{jt}^k$$

where $\varepsilon_{jt}^k \sim N(0, \sigma_{\varepsilon k}^2)$ is random noise. The troublesome household (and village) effect, α_j^k, can be written as a linear function of average household characteristics, $\bar{z}_j^k$:

$$(4\text{-}3) \qquad \alpha_j^k = \bar{z}_j^k \gamma^k + \theta^k \omega_j^k,$$

where $\omega_j^k \sim N(0,1)$, and γ^k and θ^k are parameters to be estimated. Again, this conservative econometric approach control for latent household effects by sacrificing all between-household variation in the explanatory as instruments for the latent effects. The structural parameters of interest are estimated only using the within household variation that occurs over time. Unfortunately, as with the truncated models considered in the prior section, standard fixed effect methods will not yield consistent parameter estimates and we again turn to simulation based inference following the procedures outlined in the prior section.

The results of the SML estimation of the system defined by equations (4-1) to (4-3) are presented in the second column in Table 4.[14] These SML panel estimates that control for all time invariant factors at the household (and higher) level are broadly similar to the pooled tobit estimates with the important exception that tenure security variable has become large and statistically significant. The quantitative magnitude of the is coefficient is such that it says that providing a household of immunity against administrative land reallocation would increase investment attached to land by nearly 4 hours per-*mu*, an amount equal to about 60% of the sample average investment level. For an average farm comprised of about 10 *mu*, this estimate suggests that the provision of tenure security would increase investment labor by a total of 40 hours a year for the farm in its entirety. Whatever its benefits, the current mixed system of property rights would seem to have some significant economic costs in terms of foregone investment.[15]

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[14] These estimates were estimated using 20 replications of the simulated error, ω_j.

[15] An economic evaluation of this investment would require knowledge of the marginal returns to investment as well as an estimate of how much *less* the investing household might do of other activities. The study of tenure security by Carter and Olinto (1998) finds that for capital-constrained households in Paraguay, tenure security indeed boosts investment in fixed or attached capital, but it comes at the cost of reduced investment in other types of capital.

Section 5 Conclusion: Market versus Administrative Reallocations

This paper began with the observation that while the HRS reforms in China resolved incentive problems, the left the rural economy vulnerable to increasing allocative inefficiency, especially as the non-farm economy began to boom and pull labor out of the farm sector (see Putterman, 1992). Comparing 1988 with 1993, we indeed see that the distribution of household land to labor endowments has become more disperse. Reactions to this increasing dispersion could come through either administrative reallocations that operate directly on the endowment distribution, or factor market trade that prevents the increasingly disperse distribution from spilling over and generating the sorts of allocative inefficiency that would appear as an increasingly disperse distribution of factor proportions in production. While these two mechanisms are potential substitutes for one another, they have nearly opposite implications for the post-HRS evolution of the property rights system. As Liu *et al.* (1998) argue, there is evidence that the local evolution of property rights in rural China is in fact heading off in two disparate directions, one where private rights are secured and made more marketable, and another where individual rights become increasingly restricted and subject to more regulation and reallocation.

Our efforts to use village-level panel data to gauge the forces that shape the amounts of land reallocated administratively and via the rental market have been only partially successful. While Turner *et al.* (1998) find evidence of a rational village model in which administrative land reallocations respond to both efficiency and equity concerns, we find strong evidence only of the latter. Note, however, that to the extent that access to local off-farm job opportunities in township and village enterprises is rationed in an egalitarian fashion (a possibility consistent with the labor market analysis in Yao, forthcoming), administrative reforms that maintain an egalitarian endowment distribution will also tend to preserve agricultural allocative efficiency.

Our analysis of private rental transactions was unable to satisfactorily identify the factors at the village level that shape the activeness of the rental market, and hence we have been able to put together a definitive comparison of the relative efficacy of an administrative versus a market based resolution of the post-HRS agrarian question in China. Interestingly, our comparison of the 1988 factor proportion distribution with that in 1993 shows that the latter is slightly more compressed, indicating that in the aggregate at least, rural allocative inefficiency is not increasing. In both years, the cost of allocative inefficiency is roughly estimated as 2.5% of agricultural production.

While our results make it difficult to compare the responsiveness of administrative versus market based reallocations to village level disequilibria, one clear difference between the two mechanisms is their impact on household tenure security. Panel estimates suggest that increasing household tenure security by providing households immunity against administrative reallocations boosts investment in land by 4 hours per *mu*. While seemingly modest, this figure amounts to over 50% of the average total labor time invested in land improvement and other forms of attached capital. Given the earlier evidence that administrative reallocations tend to maintain egalitarian land access (and hence function as a social safety net as Dong, 1994 and Kung, 1994 and others have argued), this finding suggests that the foregone investment and income are the implicit insurance premia paid for this social insurance embedded in the land tenure system. Whether or not this embedded insurance is reasonably priced for what it delivers, and whether it constitutes good social policy is an important question that will require additional research.

Figure 1a: Endowment Distribution, 1988-1993

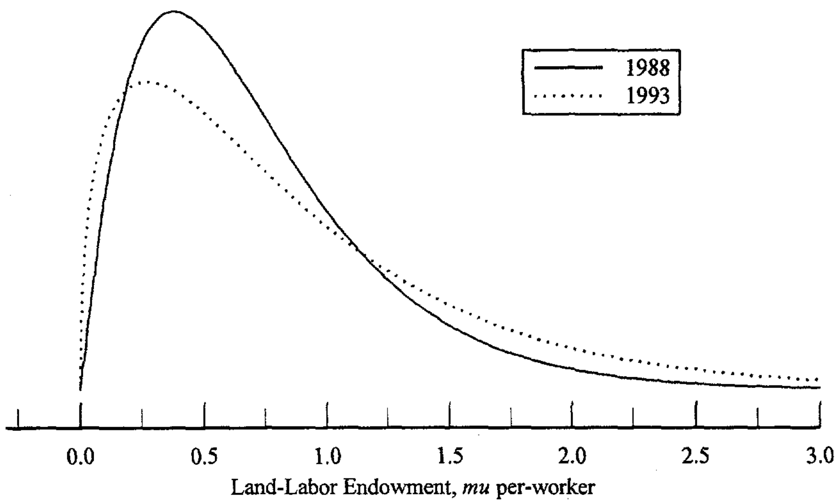

Figure 1b: Factor Intensity Distribution, 1988-1993

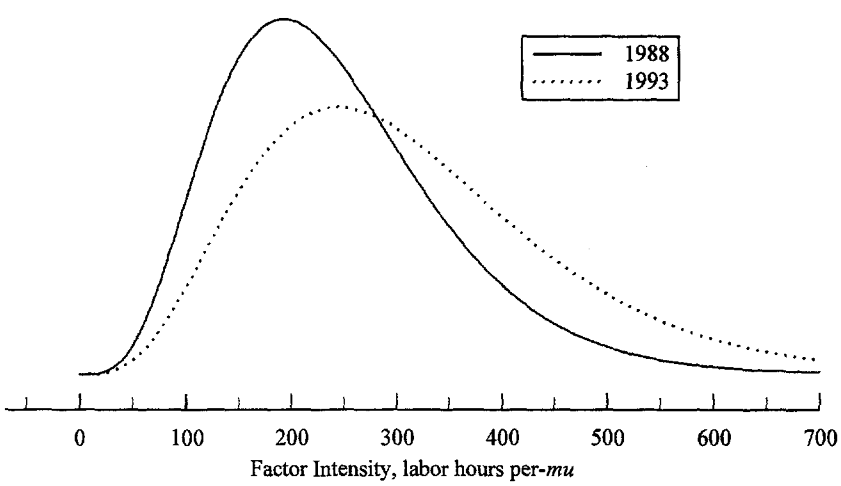

Table 1
Factor Endowments and Factor Proportions for Rice Producing Households, 1988-1993

	1988		1993
Endowment Distribution			
Mean	0.72		0.85
Coefficient of Variation			
Overall	74%		100%
Intra-Village	43%		54%
Kolmogorov-Smirnov Test for Distributional Change			
Overall		0.12* (0.6%)	
Intra-Village		0.08 (11%)	
ML Estimates of Density Function			
Shape Parameter, α			
Constant		1.46* (0.049)	
Year		-0.26* (0.062)	
Scale Parameter, β			
Constant		0.478* (0.022)	
Year		0.189* (0.04)	
Factor Proportions Distribution			
Mean	239		309
Coefficient of Variation			
Overall	46%		50%
Intra-Village	34%		32%
Kolmogorov-Smirnov Test for Distributional Change			
Overall		0.25* (0.0%)	
Intra-Village		0.04 (88%)	
ML Estimates of Density Function			
Shape Parameter, α			
Constant		2.26* (0.077)	
Year		-0.093* (0.107)	
Scale Parameter, β			
Constant		6.89* (0.248)	
Year		1.27* (0.38)	

"*" indicate that the relevant parameter or test statistic is different from zero at the 5% level.

Table 2
Village and Household Descriptive Statistics, 1988-1993

	1988		1993
Village Level Data			
% Land Rented	9.2%		8.7%
% Land Reallocated, 88-93		11.3%	
Intra-Village Coefficients of Variation			
Endowment Ratio	44%		46%
Labor Intensity	43%		40%
Non-Farm Income	93%		98%
[std dev in brackets]	[947 yuan]		[2775 yuan]
Structural Indicators			
Quotas (% expected yield)	23%		21%
Non-Farm Employment (% of total)	34%		38%
Property Rights Indicators (% villages)			
Unrestricted Rentals	73%		55%
Immune to Reallocation	65%		53%
Household Level Data			
Fixed Investment (hours/*mu*)	6.3		6.4
Endowment Ratio (*mu* per-person)	4.1		2.3
Average Household Age	32		35
Average Household Schooling Years	5.4		5.8
Fragmentation (number of parcels)	9.1		9.9
Land Quality (% flat land)	84.0		81.0

Table 3
Econometric Estimates of the
Determinants of Rental and Administrative Reallocations

	Pooled or Cross-Sectional		**SML Panel-1**		**SML Panel-2**	
	Admin Realloc	*Market Rentals*	*Admin Realloc*	*Market Rentals*	*Admin Realloc*	*Market Rental*
Structural Factors						
Constant	0.30*	-0.53**	0.28	--	0.51**	--
	(0.16)	(0.23)	(0.22)		(0.25)	
Rights Indicator, D^A, or D^R	-0.28	0.04	-0.21**	-0.26	-0.12*	-0.21
	(0.20)	(0.23)	(0.09)	(0.17)	(0.07)	(0.13)
Time Effect	--	0.15	--	0.06	--	0.11
		(0.09)		(0.06)		(0.07)
Quota	0.36	--	0.58	--	0.78**	--
	(0.51)		(0.80)		(0.26)	
D^A*Quota	-0.85	--	0.01	--	0.01	--
	(0.59)		(0.14)		(0.06)	
Dispersion Measures						
Endowment, σ^τ	-1.09**	-0.15	0.17	-0.65**	0.24**	-0.63**
	(0.36)	(0.21)	(0.12)	(0.24)	(0.10)	(0.26)
$D^{A(R)}*\sigma^\tau$	1.11**	0.49*	-0.33**	0.43	-0.61**	0.42
	(0.40)	(0.29)	(0.17)	(0.26)	(0.19)	(0.28)
Factor Intensity, σ^ℓ	-0.02	--	-0.15	--	-0.42**	--
	(0.33)		(0.13)		(0.16)	
$D^A*\sigma^\ell$	0.03	--	0.11	--	0.13	--
	(0.34)		(0.09)		(0.08)	
Off-Farm Earnings, σ^w	--	1.79	--	1.12	--	0.13
		(1.28)		(0.86)		(0.48)
$D^R*\sigma^w$	--	-1.42	--	0.01	--	-0.38
		(1.57)		(1.11)		(0.49)
Variance Terms						
σ_A or σ_R	0.04**	0.21**	0.05**	0.02*	0.05**	0.05**
	(0.01)	(0.05)	(0.01)	(0.01)	(0.01)	(0.01)
θ	--	--	--	0.17**	--	-0.04
				(0.05)		(0.09)
Δ	--	--	-0.007	--	0.05	--
			(0.04)		(0.06)	

* Significant at the 10% level.
** Significant at the 5% level.
Regional dummy variables and nuisance parameters (instruments for the α_j) not reported.

Table 4
Administrative Reallocation, Tenure Insecurity and Investment

Parameters	Pooled Tobit	SML Panel Estimates
Structural Factors		
Constant	5.1	--
	(3.4)	
Year	-1.6*	0.47
	(0.9)	(0.96)
Education	-0.18	0.88
	(0.27)	(0.73)
Age	0.12	0.08
	(0.8)	(0.13)
Fragmentation	-0.17**	-0.13*
	(0.06)	(0.08)
Land Quality	0.51	0.05
	(0.9)	(1.2)
Endowment Ratio	-0.12	0.07
	(0.15)	(0.19)
Non-Ag Employment	2.4	0.90
	(3.4)	(7.6)
Tenure security, D^A	0.45	3.9**
	(0.88)	(0.96)
Variance Terms		
θ_k		0.11**
		(1.47)
σ / σ_k^{-1}		0.11**
		(0.01)

* Significant at the 10% level.
** Significant at the 5% level.
Regional dummy variables and nuisance parameters (instruments for the α_j) not reported.

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